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Disciplines
Other Economics | Other Education | Workers' Compensation Law

WORKERS' COMPENSATION RATE REGULATION: HOW PRICE CONTROLS INCREASE COSTS*

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Abstract

In the 1980s, regulation constrained workers' compensation insurance premiums in the face of rapid growth in loss costs. We develop and test the hypothesis that rate suppression exacerbates loss growth, leading to higher losses and premiums. The empirical analysis using rating class data for eight states for the period 1985–91 confirms that rate suppression, measured by lagged residual-market share of payroll, increased loss growth. The cost-increasing effects are greater in the residual market than in the voluntary market, but premiums increased more rapidly in the voluntary market. The resulting pattern of cross subsidies between and within classes is consistent with a simple model of political influence, with subsidies to high risks and small firms at the expense of low risks and insurer equity.

I. Introduction

A fundamental prediction of economics is that price controls reduce the supply of the good or service whose price is controlled. In the case of insurance, the most visible effect of price controls is that insurers tighten underwriting standards, reducing the proportion of risks that are offered insurance voluntarily. A less obvious effect is that, since marginal insurance prices reward risk reduction by policyholders,[1] price controls may undermine incentives for loss control and hence may actually increase the losses insured against. Although price controls may thus be counterproductive and generate a deadweight loss, the positive theory of regulation[2] suggests that such negative-sum regulation may persist, at least temporarily, depending on the distribution of gains and losses.

This paper tests for these predicted allocative and distributive effects of price

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[1] See, for example, Steven Shavell, On Liability and Insurance, 13 Bell J. Econ. 120 (1982).

[2] George J. Stigler, The Theory of Economic Regulation, 2 Bell J. Econ. 3 (1971); Sam Peltzman, Toward a More General Theory of Regulation, 19 J. Law & Econ. 211 (1976); Gary S. Becker, A Theory of Competition among Pressure Groups for Political Influence, 98 Q. J. Econ. 371 (1983).

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controls in workers' compensation (WC) insurance. Employers in most states are required to purchase WC insurance to pay for medical expenses and wage loss arising out of work-related injuries and diseases or to provide proof of financial responsibility. In the 1980s, regulators in some states responded to rapidly rising loss costs by denying premium rate increases.[3] Insurers responded by reducing the proportion of risks accepted in the voluntary market, assigning the remainder to residual-market mechanisms that exist in all states. By the late 1980s, the countrywide residual-market deficit relative to premiums reached 19 percent and the residual-market share of premiums exceeded 50 percent in some states, threatening the viability of the WC insurance market. Residual-market deficits are generally financed by assessments on all WC insurers in proportion to their voluntary-market premiums in the state. These assessments may be passed forward as higher premiums to voluntary-market policyholders or passed back to insurer equity, depending on regulatory constraints and market conditions. This residual-market "tax" in turn creates incentives for policyholders to shift to self-insurance and for insurers to exit the market, thereby shrinking the tax base and increasing the tax rate on those policyholders and insurers who remain. In response to deteriorating market conditions, many states modified their statutory benefits, increased rate levels, and adopted other measures that substantially improved the situation by the mid-1990s. Nevertheless, because the underlying regulatory structure is unchanged in many states, the possibility remains for a recurrence of the 1980s meltdown in WC and in other lines of insurance that have similar regulatory structures and residual-market mechanisms.

Regulatory constraints on WC pricing may exacerbate the rate of growth of loss costs through several mechanisms. First, the suppression of rates below competitive levels in the residual market acts as a subsidy to high-risk activities. Second, to the extent that regulation reduces experience rating, employers have reduced incentives to invest in loss control and to require such investments by employees. Third, if regulation constrains the pass-through of insurer expense markups for loss control but permits the pass-through of loss costs, insurers may reduce investments in loss control, leading to increased injury rates and/or higher claim costs. Fourth, the pooling of losses among insurers through residual-market mechanisms may, under certain circumstances, reduce incentives for servicing carriers to invest in loss control.

Our empirical analysis of the effects of rate suppression on loss growth uses voluntary- and residual-market data for the 150 largest rating classes in eight states that differed in regulatory regime, for 5 years between 1985 and 1991.[4] As a measure of rate suppression, we use the lagged, class-specific residual-

[3] Workers compensation loss costs per $100 of payroll increased from $.95 in 1978 to $1.56 in 1989 and $2.4 in 1991. See Robert W. Klein, Eric C. Nordman, & Julienne L. Fritz, Market Conditions in Workers' Compensation Insurance (1993); and William J. Nelson, Workers' Compensation: Coverage, Benefits, and Costs, 1990–91, 56 Soc. Sec. Bull. 68 (1993).

[4] The 5 policy years span slightly different calendar years in different states.

market share of payroll, as an instrument for the proportion of risks for whom insurers expect the voluntary-market rates to be inadequate. We find that loss growth is positively and significantly related to rate suppression, with larger effects in the residual market than the voluntary market, as expected. By contrast, premium growth exceeds loss growth in the voluntary market, consistent with the cross-subsidy hypothesis. We show that the pattern of cross subsidies between and within classes in a state is consistent with a simple model of political influence.

Most previous studies of insurance rate regulation have focused on industry or consumer capture of the regulatory process.[5] Rate regulation has generally been measured by an indicator variable for statutory type of regulation, in particular, prior approval versus open competition. The empirical evidence has been inconclusive, possibly reflecting real variation in the way that prior-approval regulation is implemented in different states and over time.[6] Moreover, as David Appel, Michael McMurray, and Mark Mulvaney note, the form of regulation may be endogenous; specifically, high-cost states may have been more likely to deregulate.

The possibility that rate regulation may lead to higher losses is discussed but not analyzed empirically by Scott Harrington, Patricia Danzon, Orin Kramer, and Robert Klein, Eric Nordman, and Julienne Fritz.[7] Robert Kaestner and Anne Carroll and Anthony Barkume and John Ruser examine the effects of deregulation on premiums and injury rates, but the evidence is inconclusive.[8] Kaestner and Carroll conclude that regulation leads to higher premiums, which they attribute to industry capture of the regulatory process, and that high premiums reduce injury rates, owing to increased incentives for loss control. Effects of regulation on premiums are inferred from a single-year cross-section regression of wage rates on a prior-approval indicator and other controls. Such cross-sectional es-

[5] For example, David Appel, Michael McMurray, & Mark Mulvaney, An Analysis of the Net Costs of Workers' Compensation Insurance (1992); Klein, Nordman, & Fritz, *supra* note 3; Anne M. Carroll & Robert Kaestner, The Relationship between Regulation and Prices in the Workers' Compensation Insurance Market, 8 J. Reg. Econ. 149 (1995); Timothy P. Schmidle, The Impact of Insurance Pricing Deregulation on Workers' Compensation Costs, 8 Workers' Compensation Monitor 1 (1995). For a summary of the literature on automobile regulation, see Scott E. Harrington, The Impact of Rate Regulation on Automobile Insurance Loss Ratios: Some New Empirical Evidence, 3 J. Ins. Reg. 182 (1984).

[6] As of the early 1990s, very few states had true competitive rating systems that permit insurers to alter rates without close regulatory scrutiny. Moreover, regulation of residual-market rates could still effectively suppress voluntary-market rates.

[7] Scott E. Harrington, Rate Suppression (Presidential Address), 59 J. Risk & Ins. 185 (1992); Patricia M. Danzon, Administrative Costs, Price Regulation and Efficiency: A New Look at Old Issues (paper presented at the NCCI Conference on Workers' Compensation Issues, Philadelphia 1992); Orin S. Kramer, Rate Suppression, Rate-of-Return Regulation, and Solvency, 10 J. Ins. Reg. 523 (1992); and Klein, Nordman, & Fritz, *supra* note 3.

[8] Robert Kaestner & Anne M. Carroll, New Estimates of the Labor Market Effects of Workers' Compensation Insurance, 63 S. Econ. J. 635 (1997); and Anthony J. Barkume & John W. Ruser, Deregulating Property-Casualty Insurance Pricing: The Case of Workers' Compensation, in this issue, at 37.

timates are likely to be unreliable both because prior-approval regimes differ across states and because of the dynamic effects of regulation on loss costs. Specifically, if regulation raises premiums and this reduces losses, as Kaestner and Carroll hypothesize, premiums must ultimately be lower in regulated states, unless regulatory capture raises markups over loss costs to implausibly high levels. Similarly, if regulation suppresses premiums, which leads to higher losses, as we hypothesize, premiums must eventually be higher to cover the higher loss costs in regulated states, which may lead to regulatory correction as occurred in the 1990s WC reforms. Thus the estimated cross-sectional relationship between average premiums and a regulatory regime indicator depends critically on the point in the cycle at which the relationship is estimated, which may differ across states at any point in time, exacerbating the intrinsic heterogeneity within type of regulatory regime. More generally, their regulatory capture theory cannot explain the persistence and growth of residual markets over their 1983–90 period.

Barkume and Ruser, using a longer time series, find that states that dropped both prior approval and rate bureau pricing had lower injury rates and lower premiums. They conclude that their results are consistent with two very different theories of regulation. Their cartel theory predicts that regulation raises prices, thereby protecting insurers who are inefficient at loss control, which leads to higher loss costs. Their alternative theory, like ours, hypothesizes that regulation suppresses rates, thereby reducing incentives for loss control, which leads to higher injuries and premiums.

Our approach offers several advantages over these previous studies. First, we measure rate suppression directly by the size of the residual market, instead of inferring rate adequacy from a regulatory regime indicator. Barkume and Ruser concur that this is the best approach to resolve the ambiguity of their findings. Second, by estimating the effects of rate suppression (residual-market share) on loss and premium growth rates, our estimates avoid the ambiguity of cross-sectional analysis, that rate suppression tends to lower premiums but the resulting loss increase tends to raise premiums, leading to an ambiguous overall predicted relationship between regulation and premium levels in a single cross section. Third, we estimate effects separately for the residual and voluntary markets, which permits tests for differential effects of regulation on loss and premium growth in the voluntary and residual markets. Fourth, our study is the first to analyze the redistributive effects of regulation between policyholders using class-level data. A further advantage of our data is that losses, premiums, and rate suppression refer to the same class-level exposure base. Both Barkume and Ruser and Kaestner and Carroll draw premiums and injury rates from different databases, which raises the possibility that difference in sample frame may bias estimates of the differential effects of regulation on premiums and losses.

In this paper, Section II provides background on workers' compensation insurance and prior work. Section III develops our hypotheses on the cost-increasing effects of regulation. Section IV describes the data and methodology.

Section V reports the analysis of cost growth. Section VI reports the analysis of cross subsidies. Section VII concludes.

II. Background and Previous Studies

A. *Workers' Compensation Rating*

Regulation of workers' compensation insurance rates in most states historically required all insurers to use the same rates, rating classes, and experience-rating plans. Rate advisory organizations—the National Council of Compensation Insurers (NCCI) in most states—collected loss and expense data from all insurers and used these data to develop and file class-level "advisory rates" on behalf of the industry. Since the 1970s, most states have permitted individual insurers to file for deviations from the advisory rates, generally the same percent for each rating class. In the 1980s and 1990s, advisory rates were replaced by "prospective loss costs," from which insurers can file deviations and add their own expense and profit loading to derive their manual rates, by class. Some states eliminated prior-approval requirements entirely.

An individual employer's actual rate per $100 payroll is a payroll-weighted average of the manual rate for each class employed by the firm, adjusted for experience rating (the employer's own loss experience relative to the class average) and schedule-rating credits/debits based on individual underwriting assessment. Large firms are fully self-rated, medium-sized firms pay a weighted average based on their own experience and class rates, and very small firms pay pure class rates. In addition, insurers may compete through dividends to above-average-risk employers and through service quality. In the 1980s and early 1990s, increasingly stringent regulation of NCCI manual rates led to fewer downward deviations, smaller schedule-rating credits and dividends, and larger residual markets and residual-market deficits.

Residual-market mechanisms provide insurance at regulated rates to firms that cannot obtain coverage in the voluntary market. The most common residual-market structure assigns policyholders to designated servicing carriers, who service the policy without assuming underwriting risk. Residual-market deficits are usually prorated across all insurers in proportion to their voluntary-market premium in the state. The expected additional residual-market assessment thus acts as an ad valorem tax that is part of the marginal cost of writing new voluntary-market business. How far this tax can be passed through to policyholders in the voluntary market depends on the regulation and costs of self-insurance,[9] regulation of voluntary-market rates, and residual-market rates that constrain the rates that can be charged in the voluntary market. Even if insurers cannot fully recoup residual-market deficits by pass-through to the voluntary market, insurer exit

[9] Self-insured employers avoid all residual-market assessments. Switching to self-insurance is constrained by regulation, including minimum size requirements, and by switching costs for employers.

may be slow, given quasi rents on sunk investments in tangible and intangible capital.[10]

The statutory purpose of rate regulation is to assure that rates are not excessive, inadequate, or unfairly discriminatory. However, the regulatory process may be exploited by politically powerful groups to obtain rates that differ systematically from expected loss costs, conditional on available information. Rate suppression and residual-market mechanisms can generate cross subsidies within and between classes, as well as from insurer capital. This is discussed further in Section VI.

B. Previous Studies

As discussed in the Section I, several previous studies have tested the hypothesis of industry versus consumer capture of the regulatory process on average, using indicator variables for regulatory type (prior approval versus open competition, mandatory bureau rates, and so on). Results have been inconclusive, plausibly because details of structure and implementation of a given regulatory type differ across states and over time and may be endogenous and because most prior studies have ignored the feedback from regulation to losses and ultimately to premiums, which confounds cross-sectional measures of effects of regulation on premiums or loss ratios.

Several studies of the moral hazard effects of weak experience rating have confirmed that higher WC benefits tend to increase the frequency and duration of claims, particularly in firms that are only weakly experience rated.[11] Richard Butler examines the determinants of WC cost growth, focusing on increases in statutory benefits, declines in waiting periods, and demographic changes in the workforce.[12] He concludes that loss growth largely reflects increased claims reporting rather than increased employer or employee risk-taking behavior. Butler and John Worral and Alan Krueger report similar results.[13] These studies do not test for effects of rate suppression on loss growth, which we find to be significant after controlling for benefit structure changes.

There has been little analysis of cross subsidies between policyholders in workers' compensation—either measurement or modeling of the political process. One recent exception is by Wook Kwon and Martin Grace, who examine the

[10] Harrington, *supra* note 7.

[11] For example, James R. Chelius & Karen Kavanaugh, Workers' Compensation and the Level of Occupational Injuries, 55 J. Risk & Ins. 315 (1988); Bernard Fortin & Paul Lanoie, Substitution between Unemployment Insurance and Workers' Compensation: An Analysis Applied to the Risk of Workplace Accidents, 49 J. Pub. Econ. 287 (1992); John W. Ruser, Workers' Compensation Insurance, Experience Rating, and Occupational Injuries, 16 Rand J. Econ. 487 (1985); Bruce D. Meyer, W. Kip Viscusi, & David L. Durbin, Workers' Compensation and Injury Duration: Evidence from a Natural Experiment, 85 Am. Econ. Rev. 322 (1995).

[12] Richard J. Butler, Economic Determinants of Workers' Compensation Trends, 61 J. Risk & Ins. 383 (1994).

[13] Richard J. Butler & John D. Worral, Claims Reporting and Risk-Bearing Moral Hazard in Workers' Compensation Market, 58 J. Risk Ins. 191 (1991); Alan B. Krueger, Incentive Effects of Workers' Compensation Insurance, 41 J. Pub. Econ. 73 (1990).

effects of residual-market deficits, current and lagged, on premium/loss ratios (a proxy for the expense and profit loading) in the voluntary market, the residual market, and the statewide aggregate market, using state-level data.[14] They conclude that about one-fourth of residual-market assessments are passed through to the voluntary market and that current-year residual-market deficits negatively affect both residual-market and aggregate-market loadings. These findings may be biased if current residual-market deficits (and possibly loss costs) are influenced by current premium levels.

III. Theory

Workplace accidents are multiparty accidents, for which optimal prevention requires care by employers and employees. Employers are strictly liable and are required to carry liability insurance. If insurance prices cannot reflect the expected costs of risky behavior, for example, because of asymmetric information, moral hazard undermines incentives for optimal loss control.[15] Insurance price regulation may further distort marginal insurance prices, undermining loss control incentives of employers, employees, and insurers through several channels, and, hence, lead to higher loss costs.

A. Cost-Increasing Effects of Rate Suppression

Subsidies to High-Risk Behavior. Rate regulation tends to damp experience rating, by reducing the experience modification factors and/or by reducing the base rate to which the factors apply. This reduces employers' incentives to invest in loss control, including control of employees' postinjury moral hazard—overuse of rehabilitation services and delay in return to work. With pure class rating, the employer who invests in accident prevention or claim control bears all the costs, but any benefits of lower claim costs are diffused across all policyholders in the class. Suppression of experience rating is thus expected to raise injury rates and claim costs, conditional on the distribution of firms and occupations. In addition, suppression of rate differentials may lead to an increase in the proportion of high-risk occupations and firms. These effects are expected to be greater in the residual market than in the voluntary market.

Constraints on Insurer Expense and Profit Loadings. Regulation may constrain an insurer's expense and profit markup but (largely) permit the pass-through of an increase in expected losses. If so, insurers may reduce loss control even if total costs thereby increase, because the resulting increase in losses can be (largely) passed through as higher premiums whereas the loss control expense

[14] Wook J. Kwon & Martin F. Grace, Examination of Cross Subsidies in the Workers' Compensation Market, 15 J. Ins. Reg. 256 (1996).

[15] Shavell, *supra* note 1.

comes out of profit.[16] Reduction in insurer injury prevention programs could increase the frequency and severity of injuries; in addition, diminished effort to control invalid or excessive claims could increase claim frequency and severity, conditional on injuries. On the other hand, profit constraints that reduce quality competition could lead insurers to deny more claims, which could reduce reported claim frequency, ceteris paribus. Thus the predicted effect of insurer response to rate suppression on reported claim frequency is ambiguous, particularly for the voluntary market where quality competition would exist in the absence of rate suppression.

Pooling of Residual-Market Deficits. In the dominant residual-market model during our sample period, one or more servicing carriers are selected to administer the program in return for a fixed proportion of premiums. Any operating deficit (excess of losses plus expenses over premiums) is prorated to all insurers in proportion to their voluntary WC premium volume in the state. Since the servicing carrier bears all the costs of loss control but receives only a fraction of any resulting reduction in losses, incentives for loss control may be suboptimal. This effect would exacerbate loss growth in the residual market relative to the voluntary market, ceteris paribus. However, this free-rider incentive may be countered by the monitoring and incentive programs that exist in most states, the insurer's concern for reputation and future business, and the incentives that the servicing carrier faces for its voluntary-market business, where it bears all costs. Separate loss control programs could be established for its residual- and voluntary-market business, but this would entail fixed costs. If so, any cost-increasing effects from pooling residual-market deficits are predicted to increase with the size of the residual market.

B. Winners and Losers from Rate Suppression

Modeling the market for influence over workers' compensation rates requires assumptions about the incidence of premiums and service quality. Simple theory and empirical evidence conclude that the long-run incidence of workers' compensation premiums is largely on workers, through compensating wage offsets.[17] Thus the primary beneficiaries of rate suppression are workers for whom rates are suppressed below competitive levels, who benefit from higher wages or more jobs than would be available if insurers charged actuarial rates. Some benefit may also accrue to owners of immobile factors such as entrepreneurial capital,

[16] Optimal loss control requires investing to the point where the marginal dollar invested in prevention or claims management reduces expected loss cost by just 1 dollar.

[17] For example, W. Kip Viscusi & Michael J. Moore, Workers' Compensation: Wage Effects, Benefit Inadequacies, and the Value of Health Losses, 69 Rev. Econ. Stat. 249 (1987); Jonathan Gruber & Alan B. Krueger, The Incidence of Mandated Employer Provided Insurance: Lessons from Workers' Compensation Insurance (working paper, Princeton Univ.. Indus. Rel. Section, November 1990). For effects of unionization on wage offsets and structure of risk compensation, see Price V. Fishback, Review of Fatal Tradeoffs: Public and Private Responsibilities for Risk, 32 J. Econ. Literature 131 (1992).

if insurance rates deviate from levels that were anticipated when wage rates were set or from levels that prevail for similar employees in other states.[18]

The costs of rate suppression—higher loss costs and reduced quality—may be partly borne by employers and employees in the cross-subsidized firms, through lower dividends, lower experience-rating credits, and reduced service quality, including denial of or delay in claim payment. Denial of medical claims by the WC insurer may shift costs to the employers' first-party medical insurance, which is typically self-insured for medium and large employers.

But if persistent rate suppression is a politically rational strategy for some classes, some of the resulting cost increase must be shifted to other policyholders or insurer equity. The results by Kwon and Grace suggest that both forms of shifting occur.[19] Here we focus on cross subsidies between policyholders within a state. Between-class transfers can result from suppression of class-rate relativities for manual rates. Within-class transfers from low to high risks can result from suppression of relative rates for residual and voluntary business, from regulatory constraints on experience rating and ex post dividends, and from other adjustments that reward policyholders for good loss experience. A more complex model and more comprehensive data would be required to analyze the shifting of losses to insurer equity.

An alternative hypothesis to explain the observed phenomena is regulatory lag in adjusting premiums, following an exogenous increase in losses.[20] These hypotheses are not necessarily mutually exclusive, since political pressure may contribute to the lag in premium adjustment. While some of our tests cannot distinguish between a purely mechanical regulatory lag model and one exacerbated by political influence, a mechanical regulatory lag model alone cannot explain the evidence on redistribution between and within classes, as discussed further below.

C. Testable Implications

The key prediction of this analysis is that rate suppression increases loss growth. With the available data, we cannot distinguish effects on injuries versus claims or the impact of incentive distortions to employers, employees, and insurers. However, the theory predicts larger effects in the residual market, owing to both greater distortion of employer/employee incentives and possible pooling effects for insurers. A possible exception is for claim frequency, where the employer/employee moral hazard that leads to higher injuries and claims could

[18] Patricia M. Danzon, The Political Economy of Workers' Compensation: Lessons for Product Liability, 78 Am. Econ. Rev. 305 (1988), uses a general equilibrium model to show that the incidence of state-mandated workers' compensation or health insurance may be partly on immobile factors, such as entrepreneurial capital in small firms.

[19] Kwon & Grace, *supra* note 14.

[20] For example, benefit costs per $100 of payroll increased at an average annual rate of 4.2 percent for 1978–84 and 6.2 percent for 1984–89. Klein, Nordman, & Fritz, *supra* note 3.

be offset by higher claim rejection rates by insurers due to reduced incentives to compete on service quality, leading to lower overall reported claim frequency in residual markets. Cross subsidies are expected to result in transfer from low risks to high risks. Other factors that influence cross subsidies are discussed further below.

IV. Methodology and Data

A. Model

The theory outlined suggests the following basic estimating equation:

$$[Y^t_{ijs}/Y^{t-1}_{ijs} - 1] = \beta_1 R^{t-1}_{ijs} + \beta_2 (R^t_{ijs}/R^{t-1}_{ijs}) + \beta_3 B^{t-1}_s + \varepsilon_j + u^t_{ijs}, \qquad (1)$$

where $[Y^t/Y^{t-1} - 1]$ is percentage growth in loss per \$100 payroll between period t and period $t-1$; subscript i denotes class, subscript j denotes industry, and subscript s denotes state; R^{t-1} is lagged residual-market share (a measure of rate suppression); B is a vector of state fixed effects or other state-specific characteristics that affect loss growth, such as statutory benefit change; R^t_{ijs}/R^{t-1}_{ijs} is a control for selection bias (see below); ε_j is a vector of fixed industry effects; and u^t_{ijs} is a stochastic error term. We estimate equation (1) for the voluntary and residual markets separately. The hypothesis that rate suppression exacerbates loss growth predicts that $\beta_1 > 0$, with greater effect in the residual market than the voluntary market.

B. Data

The data are from the NCCI for 5 consecutive policy years per state between 1985 and 1991 and are reported by individual rating class. For most of the analysis, we use data for eight states (Alabama, Florida, Georgia, Illinois, Louisiana, Maine, Michigan, and Virginia) that represent different regulatory regimes. These eight-state data are for the 150 largest classes in each state, with voluntary-market and residual-market experience reported separately for each class. Although these 150 classes represent only roughly one-third of the classes in each state, they account for over 80 percent of the market as measured by risks, premium, payroll, or losses.[21] For a larger sample of 27 states, we have classwide data, that is, data not reported separately for the voluntary and residual markets. This larger sample of classwide data is used for analysis of between-class cross subsidies. Since WC claims accrue for several years after the policy is written, for each policy year we use the forecast (“fully developed”) value of claims and losses based on the most recent estimate available.[22] We convert all dollar values

[21] Sixty-four of the classes are in all eight states, 25 are in seven states, and 120 are present in only one state.

[22] Incurred losses are reported as of 1992 for all policy years. The NCCI calculates development factors based on prior experience which are used to adjust early loss reports to estimated ultimate value. The development factors are uniform for all classes in a state, hence do not reflect any differences in payout tails between classes. Provided that any resulting error is uncorrelated with our explanatory variables, our estimates should be unbiased.

to constant 1992 dollars using a wage index for payroll and indemnity losses, a medical index for medical losses, and a weighted average of the medical and wage index for premiums.

The NCCI rating classes are defined roughly along occupational lines. The class data on premiums, losses, payroll, and so on, reflect the pooled experience for that class from all employers who employ the class in the state, from all insurers. The premium rate for a specific employer ("risk") would be calculated as the payroll-weighted average of rates for the classes represented by its employees, but these class-level data do not report premiums for individual employers or insurers. The data do report the number of employers for which that class is the dominant share of payroll (Risks) and the number of employers reporting any data for that class (Classcount). Manual and standard (after experience rating) premium rates are expressed per \$100 of payroll.

The loss data at the class level exhibit extreme stochastic variation across years; payroll and number of risks are also unstable, presumably because of switching to self-insurance and employment changes within insured firms. Because the annual time trends are very unstable, we aggregate the most recent 3 years as period t and the earliest 2 years as period $t - 1$ and define loss growth as the change between t and $t - 1$.[23] Using multiyear mean values averages out some of the stochastic variation.

Loss Growth. We use three measures of loss growth: absolute loss growth, simple percentage loss growth, and the natural logarithm of percentage loss growth.[24] These different measures could yield different conclusions. For example, it is possible that classes with relatively high initial losses could experience more rapid absolute loss growth than classes with a lower initial loss level, but the high-cost class could still have a lower percentage growth rate. For the log measure, the growth in total loss is equal to the sum of the growth in claim frequency and severity.

Rate Suppression. Our primary measure of rate suppression is the lagged class-specific percentage of payroll in the residual market.[25] This is the best available class-specific measure of the percent of employees for whom voluntary-market rates are perceived by insurers to be inadequate, owing to either suppression of voluntary-market rates or suppression of residual-market rates, which then constrain the rates that can be charged in the voluntary market. We use the lagged value of the residual-market share as an instrument for contemporaneous residual-market share, which could be endogenous. To the extent that the lagged residual-market share is an imprecise measure of contemporaneous rate inadequacy, the resulting measurement error reduces the precision of coefficient es-

[23] The 3-year mean loss per \$100 payroll is the sum of losses divided by the sum of payroll, which is effectively a weighted mean.

[24] Let Y_t denote the mean for the most recent 3 policy years and Y_{t-1} denote the mean for the prior 2 years. Absolute loss growth is $Y_t - Y_{t-1}$, percentage loss growth is $(Y_t/Y_{t-1} - 1)$, and the natural log of percentage loss growth is $\ln(Y_t/Y_{t-1})$.

[25] The lagged value is the mean over the earliest 2 years.

timates, leading to downward-biased estimates of statistical significance. As an alternative measure of regulatory stringency, we also use the ratio of filed to approved rates at the statewide level.[26] This measure of mean statewide rate inadequacy may be imprecise for individual classes.

Other Variables. To control for state-specific statutory benefit changes, which may trigger exogenous loss shocks, we include the NCCI's estimate (lagged) of predicted percentage change in loss costs for the state. The estimated effects of this benefit change variable on loss growth may be downward biased, if states with rapidly growing losses and large residual markets are more likely to reduce their statutory benefits.[27] To test the hypothesis that the insurer moral hazard effect of residual-market risk pooling increases with the size of the residual-market deficit, we include the lagged residual-market deficit per $100 payroll in the state.[28] A vector of fixed industry effects is included to control for industry-specific determinants of loss growth.[29] In some specifications, we substitute a vector of fixed state effects for the substantive state-level variables, to control for other, unmeasured state characteristics that influence loss growth. The use of class-level data should reduce the risk of bias due to unobserved state-specific factors.

Outliers. We exclude Maine from some of the analysis reported here. Its abnormally large residual-market size and unique residual-market mechanism, with deficit sharing by employees, make its experience atypical, which could bias conclusions for other states. Including Maine generally yields similar but less precise coefficient estimates. We exclude from the regressions observations with missing data on premiums or payroll and extreme outliers, defined as observations with any loss growth measure (absolute, relative, and percentage growth in total losses, claim frequency, or severity) more than 4 standard deviations from the mean for that variable. Such extreme deviations most likely reflect the stochastic variation in losses for small classes. This reduced the sample from 1,050 (after omitting Maine) to 860 observations.

Selection Bias. The theoretical model considers rate suppression's effect on loss growth for a group of policies with specified risk characteristics. The observed measures of loss growth in the residual market may be downward (upward) biased when the residual market is growing (shrinking), assuming that the highest risks are assigned first to the residual market, such that the expected cost

[26] For previous uses of this measures, see Klein, Nordman, & Fritz, *supra* note 3; and Scott E. Harrington & Patricia M. Danzon, Rate Regulation, Safety Incentives and Loss Growth in Workers' Compensation Insurance, 73 J. Bus. 569 (2000).

[27] The NCCI measure is a weighted average of estimated effects of statutory changes on the severity of different types of claims, using the most current distribution of injury costs, that is, without adjustment for potential changes in accident or claim frequency.

[28] The correlation between statewide residual-market deficit and class-specific residual-market share is .45 (excluding Maine).

[29] We use the NCCI class and industry definitions, which are similar but not identical to Standard Industrial Classification definitions.

of the marginal risk is less than the average risk. A similar downward bias in estimated loss growth occurs in the voluntary market because the average expected loss per exposure unit declines as the voluntary market shrinks, assuming that the lowest risks are the last to be assigned to the residual market. We therefore include the growth in the residual-market share, as a control for selection bias. If shifting of risks between the voluntary and residual markets is significant, including this control is expected to increase the estimated effect of rate suppression on loss growth in the voluntary and residual markets separately. Estimates of classwide loss growth, for voluntary and residual markets combined, should be unaffected by switching between the submarkets, assuming that the total exposure base remains unchanged.

The self-selection of lower risks to self-insurance could lead to upward-biased estimates of regulation-induced loss growth, if residual-market rate suppression raises the deficit tax on voluntary-market premiums and hence increases incentives for low risks to self-insure.[30] However, our measure of rate suppression—residual-market share of payroll—primarily reflects rate suppression in the voluntary market, which should reduce the cost of market insurance relative to self-insurance and hence reduce incentives to self-insure, ceteris paribus. Any selection bias due to self-insurance is more likely to affect the coefficient estimates on the statewide residual-market deficit, which measures the potential tax on voluntary-market premiums.

Heteroscedasticity. Since the dependent variable reflects loss experience for a class of risks, the residual variance σ_{ijs} may be related to size of insured payroll, via the law of large numbers, but may also be related to other unmeasured factors. A White test indicates significant heteroscedasticity for some ordinary least squares (OLS) regressions. We therefore report approximate t-statistics using White standard errors. Preliminary analysis showed little if any advantage from using weighted least squares (WLS), with weights defined as either the square root of payroll or the predicted value from a regression of OLS squared residuals on the inverse of payroll and number of risks.

C. Descriptive Statistics

Table 1 reports alternative measures of mean statewide rate suppression: the residual-market share of premiums, payroll, and risks; the loss ratio, often used as a measure of price to the insured or profitability to the insurer;[31] and the ratio of filed to approved rates. The mean residual-market share of payroll is similar

[30] Anne M. Carroll, The Role of Regulation in the Demand for Workers' Compensation Self-Insurance, 13 J. Ins. Reg. 168 (1994), provides evidence that higher ratios of premiums to losses increase the self-insurance proportion, implying that overall rate suppression might reduce the use of self-insurance. No relationship is found between residual-market share and the self-insurance variable.

[31] The statewide loss ratio is the unweighted mean of class-specific loss ratios, where each class loss ratio is the 5-year mean of losses divided by the 5-year mean of premiums.

TABLE 1

MEASURES OF RATE SUPRESSION, BY STATE 5-YEAR MEANS, 1985–91

	Share of the Residual Market								Filed/Approved					
	Payroll		Risks		Premium		Loss Ratio		Voluntary		Residual		Total	
State	Value	Rank	Value	Rank	Value	Rank	Value	Rank	Value	Rank	Value	Rank	Value	Rank
Michigan	.060	1	.124	1	.076	1	.828	2	1.007	2	1.013	1	1.008	1
Georgia	.111	2	.248	2	.158	2	.922	4	1.080	4	1.699	7	1.180	5
Illinois	.115	3	.260	3	.141	3	.681	1	1.000	1	1.167	4	1.027	2
Virginia	.135	4	.289	4	.149	4	.916	3	1.201	6	1.210	5	1.203	6
Alabama	.212	5	.387	5	.228	5	1.139	6	1.065	3	1.060	3	1.064	3
Florida	.220	6	.513	6	.208	6	1.381	8	1.138	5	1.050	2	1.118	4
Louisiana	.364	7	.562	7	.397	7	1.053	5	1.322	7	1.350	6	1.339	7
Maine	.775	8	.869	8	.787	8	1.306	7	1.569	8	1.863	8	1.771	8

TABLE 2

Characteristics of Voluntary and Residual Markets: Unweighted Means across Classes

A. 1988–91

	Claim Frequency (Claims/$10,000 Payroll)		Claim Severity (Loss/Claim)		Risks/$10,000 Payroll		Loss/$100 Payroll	
State	Voluntary	Residual	Voluntary	Residual	Voluntary	Residual	Voluntary	Residual
Michigan	.118	.199**	5,621	6,102*	.042	.100**	6.35	10.69**
Georgia	.122	.198**	3,987	5,205**	.028	.092**	5.05	10.47**
Illinois	.091	.141**	5,566	6,401**	.027	.078**	5.05	8.92**
Virginia	.077	.113**	4,549	7,857**	.034	.085**	3.62	9.07**
Alabama	.110	.183**	5,991	6,212	.030	.067**	6.19	11.10**
Florida	.110	.148**	9,295	15,352**	.035	.164**	10.51	22.60**
Louisiana	.095	.137**	7,073	7,372	.033	.086**	6.86	10.58**
Maine	.233	.145	8,775	9,466	.054	.065	16.70	13.48

B. 1985–87

	Claim Frequency (Claims/$10,000 Payroll)		Claim Severity (Loss/Claim)		Risks/$10,000 Payroll		Loss/$100 Payroll	
State	Voluntary	Residual	Voluntary	Residual	Voluntary	Residual	Voluntary	Residual
Michigan	.124	.209**	4,570	6,527**	.046	.103**	5.28	9.18**
Georgia	.144	.191**	2,827	4,139**	.033	.073**	4.08	8.07**
Illinois	.105	.146**	4,100	4,947**	.030	.076**	4.31	7.24**
Virginia	.099	.139**	3,637	4,849**	.035	.108**	3.67	6.36**
Alabama	.129	.195**	4,594	4,169	.033	.074**	5.91	7.73**
Florida	.128	.164**	7,784	10,877**	.038	.146**	10.30	17.62**
Louisiana	.112	.145**	6,094	6,723	.043	.104**	7.13	10.13**
Maine	.184	.173	9,982	7,183**	.048	.072**	16.59	12.18*

* Difference in means significant at $p = .05$.
** Difference in means significant at $p = .01$.

to the share of premium but consistently less than share of risks. This implies that firms in the residual market on average have fewer employees and/or lower wages than firms in the voluntary market. The ranking of states by loss ratios is similar to the ranking by residual-market shares, which confirms that the size of residual market is a measure of rate inadequacy. The ranking by the filed-to-approved ratio is less consistent, possibly because this regulatory stringency measure reflects the magnitude of the difference between filed and approved rates, rather than the percent of payroll for which the filed rates are inadequate.

Table 2 compares mean claim frequency, claim severity, and total loss per dollar of payroll, for the voluntary and residual markets in each state, for the base period (1985–87) and the most recent period (1988–91). The market mean is the unweighted mean across classes, and each class observation is the weighted mean over the years included.[32] In all states except Maine, both claim frequency

[32] For example, claim severity is $\sum_{t=1}^{3}\text{Loss}^t/\sum_{t=1}^{3}\text{Claims}^t$.

and total loss per $100 of payroll are significantly higher in the residual market than in the voluntary market. Claim severity is also higher in the residual market than in the voluntary market, except in Alabama, Louisiana, and Maine. The number of risks per $10,000 payroll is two to three times higher in the residual market than in the voluntary market (except in Maine), which is further evidence suggesting that residual-market risks are on average smaller firms.[33] Maine is anomalous in having the largest residual-market share (78 percent of payroll, 79 percent of premium, and 87 percent of risks), the highest loss per $100 payroll, and higher claim frequency and severity in the voluntary market than in the residual market. This may reflect random noise, given the small exposure base in the voluntary market.

The evidence in Tables 1 and 2 shows persistent differences across states in several measures of rate suppression. This is consistent with the hypothesis that deviations from actuarial rates reflect political influence rather than simply stochastic shocks, in which case they should average out over time. Persistence in rate suppression clearly spans several rating periods with clear evidence of rate inadequacy, which seems inconsistent with a purely mechanical regulatory lag model.

Table 3 reports the absolute, percentage, and log percentage measures of mean growth rates over all classes in the seven states excluding Maine. The standard deviations are often several times larger than the means because of small class size. The mean percentage growth in total loss per $100 payroll between the two time periods was 16.9 percent, or roughly 7 percent a year. Growth was over twice as high in the residual market (37 percent) as in the voluntary market (15.2 percent). This growth in losses is driven primarily by growth in claim severity, for which the marketwide average mean growth is 15.6 percent, composed of 33.9 percent for the residual market and 17.2 percent for the voluntary market.[34] Claim frequency grew 2.7 percent overall, composed of an 8.5 percent increase for the residual market and only a .7 percent increase in the voluntary market. The more rapid growth in claim frequency in the residual market than in the voluntary market suggests that any incentive for insurers to deny claims in response to rate suppression is more than offset by the claim-increasing effects of weak incentives for loss control by employers, employees, and insurers. For claim severity, which has higher stochastic variance, the ranking of the residual and voluntary markets depends on the measure used.

The mean increase in premium per $100 of payroll is 20.5 percent for the total market, composed of 23.4 percent in the residual market and 19.6 percent in the voluntary market. This compares to an increase in loss per $100 payroll of 37.1 percent for the residual market and 19.6 for the voluntary market. This

[33] Number of risks per $10,000 payroll is an approximate measure of firm size because risks includes only the firms for which this is the dominant class, whereas payroll includes the payroll of all firms with this class.

[34] It is possible for the mean for the total market to be less than the means for the two submarkets.

TABLE 3

LOSS GROWTH MEANS AND STANDARD DEVIATIONS BY TOTAL, RESIDUAL, AND VOLUNTARY MARKETS

Variable and Market	Mean	SD
Absolute Growth ($Y^t - Y^{t-1}$):		
Loss/Payroll:		
Total	.0069	.0273
Residual	.0105	.0810
Voluntary	.0041	.0297
Loss/Claim (Severity):		
Total	.5350	2.2793
Residual	.3516	6.2751
Voluntary	.4678	2.5831
Claims/Payroll (Frequency):		
Total	.0488	2.1546
Residual	.4673	5.5275
Voluntary	(.1972)	2.3931
Premium/Payroll:		
Total	.0123	.0161
Residual	.0153	.0209
Voluntary	.0114	.0156
Percentage Growth ($Y^t/Y^{t-1} - 1$):		
Loss/Payroll:		
Total	.1689	.3779
Residual	.3705	.8801
Voluntary	.1524	.4247
Loss/Claim (Severity):		
Total	.1563	.3666
Residual	.3385	.9321
Voluntary	.1724	.4389
Claims/Payroll (Frequency):		
Total	.0265	.2110
Residual	.0847	.3881
Voluntary	.0067	.2436
Premium/Payroll:		
Total	.2049	.1961
Residual	.2343	.2370
Voluntary	.1963	.1968
log Percentage Growth ($ln(Y^t/Y^{t-1})$):		
Loss/Payroll:		
Total	.1038	.3292
Residual	.0996	.7201
Voluntary	.0729	.3822
Loss/Claim (Severity):		
Total	.0978	.3084
Residual	.0751	.6970
Voluntary	.0949	.3591
Claims/Payroll (Frequency):		
Total	.0060	.2009
Residual	.0245	.3397
Voluntary	(.0220)	.2419
Premium/Payroll:		
Total	.1732	.0161
Residual	.1916	.0209
Voluntary	.1659	.0156
Other:		
Total: Deficit/Payroll ($t - 1$)	.1700	.1000
Total: Benefit Growth ($t - 1$)	1.0206	.1328
Total: Class Residual-Market Share ($t - 1$)	.1573	.0269
Total: Residual-Market Share Growth	.0332	.0716

suggests an increase in cross subsidies from the voluntary market to the residual market, which is confirmed in the multivariate analysis below. The mean residual-market share of payroll is 20.4 percent including Maine and 15.7 percent excluding Maine. The growth in statewide residual-market share is 4.1 percent including Maine and 3.3 percent excluding Maine. The inclusion of Maine doubles the estimate of residual-market deficit per $100 statewide payroll, from $.17 without Maine to $.36 with Maine.

The correlation between absolute growth and simple percentage growth in loss per $100 payroll is .73. Classes with a relatively high initial residual-market share experienced more rapid growth in their residual-market share ($r = .25$). Statutory benefit growth is negatively correlated with initial residual-market share ($r = -.13$), which suggests that benefit constraint was one response to residual-market problems.

V. Effects of Rate Suppression on Cost Growth

Tables 4 and 5 report three specifications for the effects of rate suppression on absolute and log percentage growth of losses and premiums, respectively. To control for state-specific influences on loss growth, equation (1) includes fixed state effects. Equations (2) and (3) replace the fixed state effects with substantive state variables: the lagged predicted growth in benefit costs for the state; the growth in class-specific residual-market share, to control for between-market selection effects; and the residual-market deficit per $100 payroll. The reduction in the adjusted R^2, relative to equation (1), indicates that the substantive state-level variables only partially control for all state-specific influences on loss growth. All equations include fixed industry effects.

A. Total Loss per $100 Payroll

Table 4 confirms that absolute growth in loss per $100 payroll is positively related to rate suppression. The estimated effect is almost 10 times greater in the residual market than in the voluntary market, where it is not significantly different from zero. Controlling for the statewide residual-market growth (equation (2)) boosts the size and the significance of the estimated effect of rate suppression on loss growth in the residual market, but the effect in the voluntary market is still insignificant. The coefficient of residual-market growth is negative and significant for the residual and voluntary markets separately, but it is insignificant for the total market. These findings are consistent with the selection hypothesis, that residual-market growth lowers the mean risk level in both the residual and voluntary submarkets. Thus equation (1), which does not control for selection bias, provides downward-biased estimates of the effects of rate suppression on loss growth. The results for log percentage growth (Table 5) are similar, implying that rate suppression resulted in greater percentage cost growth for classes that started at an initially higher level of costs.

For the residual market, these effects are economically as well as statistically

TABLE 4

EFFECT OF RATE SUPPRESSION ON ABSOLUTE LOSS GROWTH

	LOSS/PAYROLL			CLAIM SEVERITY			CLAIM FREQUENCY			PREMIUM/PAYROLL		
	Total	Residual	Voluntary	Total	Residual	Voluntary	Total	Residual	Voluntary	Total	Residual	Voluntary
Equation (1)[a]:												
Residual Market												
Share$_{is}$ $(t-1)$	.0160	.0516	.0065	.8911	1.9693	1.8920	−.3727	3.6227	−1.3695	.0303	.0333	.0281
	(1.630)	(2.103)	(.586)	(1.027)	(1.046)	(1.676)	(.583)	(2.322)	(.056)	(6.469)	(5.705)	(5.201)
Adjusted R^2	.0468	.0234	.0244	.0371	.0186	.0164	.1168	.0468	.1124	.3434	.3098	.3264
Equation (2)[a]:												
Residual Market												
Share $(t-1)$	.0205	.0733	.0083	1.2660	3.8926	1.7929	−.4243	2.4826	−1.2820	.0464	.0517	.0439
	(2.184)	(3.188)	(.837)	(1.612)	(2.222)	(1.836)	(.726)	(1.782)	(1.548)	(8.998)	(8.192)	(7.431)
Residual Market												
Share Growth$_{is}$	−.0117	−.0960	−.0469	−3.0004	−8.7733	−3.0262	3.3557	−1.0842	−.1302	−.0116	−.0412	−.0144
	(.763)	(2.184)	(2.821)	(2.216)	(2.464)	(1.912)	(2.956)	(.468)	(.099)	(1.593)	(4.858)	(1.954)
Benefit Growth$_s$ $(t-1)$	.0444	.0003	.0499	3.6169	−8.0533	5.3713	−3.6995	29.9557	−7.9992	.0992	.2348	.0718
	(1.412)	(.003)	(1.377)	(1.595)	(.998)	(2.106)	(1.259)	(3.580)	(2.554)	(6.246)	(8.836)	(4.827)
Adjusted R^2	.0370	.0210	.0199	.0342	.0126	.0175	.0634	.0244	.0661	.2363	.2306	.2140
Equation (3)[b]:												
Residual Market												
Share$_{is}$ $(t-1)$	.0164	.0593	.0063	1.0522	2.5722	2.0464	−.5525	3.3728	−1.6315	.0308	.0370	.0283
	(1.699)	(2.448)	(.584)	(1.236)	(1.372)	(1.893)	(.847)	(2.174)	(1.837)	(6.194)	(5.823)	(4.912)
Deficit/Payroll$_s$ $(t-1)$	.0045	−.0015	−.0101	−.4742	−.0543	−1.3765	1.2337	−2.0243	.6310	.0265	.0160	.0257
	(.612)	(.065)	(1.256)	(.695)	(.030)	(1.784)	(2.129)	(1.525)	(.992)	(7.480)	(3.663)	(7.050)
Benefit Growth$_s$ $(t-1)$	.0425	−.0203	.0388	2.9230	−9.9307	4.5692	−2.8434	29.4941	−7.9552	.0997	.2279	.0717
	(1.359)	(.170)	(1.074)	(1.294)	(1.232)	(1.794)	(.976)	(3.552)	(2.562)	(6.220)	(8.481)	(4.794)
Adjusted R^2	.0366	.0144	.0102	.0268	.0034	.0160	.0576	.0265	.0672	.2806	.2224	.2566

NOTE.—Absolute t-statistics using White standard errors are in parentheses.
[a] Equation includes state and industry fixed effects.
[b] Equation includes industry fixed effects.

TABLE 5

EFFECT OF RATE SUPPRESSION ON PERCENTAGE LOSS GROWTH

	Total Loss/Payroll			Claim Severity			Claim Frequency			Premium/Payroll		
	Total	Residual	Voluntary	Total	Residual	Voluntary	Total	Residual	Voluntary	Total	Residual	Voluntary
Equation (1)[a]:												
Class Residual												
Market Share$_{is}$ $(t-1)$	.0464	.3005	.0083	.0837	.0278	.1773	−.0374	.2727	−.1690	.1615	.1631	.1676
	(.427)	(1.448)	(.063)	(.751)	(.140)	(1.252)	(.633)	(2.721)	(1.763)	(3.659)	(3.294)	(3.081)
Adjusted R^2	.0452	.0101	.0577	.0290	.0045	.0329	.1275	.0446	.1179	.3463	.3214	.3175
Equation (2)[b]:												
Class Residual												
Market Share$_{is}$ $(t-1)$	.06513	.4362	.0162	.0897	.2022	.1556	−.0246	.2340	−.1393	.3912	.4181	.3908
	(.660)	(2.324)	(.140)	(.905)	(1.118)	(1.255)	(.467)	(2.673)	(1.618)	(8.094)	(7.614)	(6.671)
Residual Market												
Share Growth$_{is}$	−.1746	−.5745	−.7968	−.4172	−.4346	−.5905	.2426	−.1399	−.2063	.0912	−.1836	.0489
	(1.020)	(1.714)	(3.883)	(2.523)	(1.371)	(2.986)	(2.230)	(.830)	(1.460)	(1.260)	(2.270)	(.647)
Benefit Growth$_s$ $(t-1)$	.6427	1.2655	.6160	.8205	−.1788	1.0459	−.1779	1.4443	−.4298	1.7664	2.6430	1.4800
	(1.565)	(1.320)	(1.358)	(2.099)	(.194)	(2.449)	(.726)	(3.162)	(1.583)	(9.303)	(10.711)	(7.618)
Adjusted R^2	.0187	.0049	.0350	.0223	.0007	.0253	.0597	.0223	.0608	.1734	.1766	.1473
Equation (3)[b]:												
Residual Market												
Share$_{is}$ $(t-1)$	.0766	.3899	.0453	.1276	.1212	.2321	−.0510	.2687	−.1868	.1854	.2267	.1859
	(.711)	(1.905)	(.349)	(1.158)	(.620)	(1.704)	(.862)	(2.746)	(1.958)	(4.023)	(4.232)	(3.256)
Deficit/Payroll$_s$ $(t-1)$	−.0734	−.0805	−.2902	−.1954	.0272	−.3204	.1220	−.1076	.0302	.4209	.3125	.4069
	(.815)	(.434)	(2.745)	(2.262)	(.154)	(3.219)	(2.180)	(1.187)	(.439)	(12.072)	(7.693)	(10.756)
Benefit Growth$_s$ $(t-1)$	.5971	1.1338	.4131	.7093	−.2684	.8835	−.1123	1.4022	−.4704	1.8338	2.6394	1.5367
	(1.469)	(1.191)	(.917)	(1.828)	(.293)	(2.080)	(.463)	(3.090)	(1.748)	(9.592)	(10.556)	(7.896)
Adjusted R^2	.0182	.0021	.0245	.0206	−.0011	.0263	.0592	.0233	.0576	.2875	.2162	.2540

NOTE.—Absolute t-statistics using White standard errors are in parentheses.

[a] Equation includes state and industry fixed effects.

[b] Equation includes industry fixed effects.

significant. From Table 4, equation (2), a 10 percentage point increase in residual-market share implies an increase in absolute loss growth for the residual market of .007, which is two-thirds as large as the mean absolute loss growth of .0105 (see Table 3) for the residual market during the sample period. From Table 5, equation (2), a 10 percentage point increase in residual-market share implies a 4.4 percentage point increase in log percentage loss growth, compared to the sample mean log percentage loss growth of 10 percent.

B. Claim Severity (*Loss/Claim*)

Absolute growth in claim severity is positively related to rate suppression, with larger effects in the residual market than in the voluntary market, after controlling for selection bias. This is consistent with the hypothesis of greater employer moral hazard in the residual market and possibly suboptimal insurer loss control due to the pooling of losses and weak incentives for quality competition. The positive association between rate suppression and claim severity growth in the voluntary market plausibly reflects attenuation of experience rating and constraints on insurer expense margins that often accompany suppression of average rate levels.[35] For log percentage growth, rate suppression is not significant. This suggests that the positive effect of rate suppression on absolute severity growth is not large enough to generate a positive percentage effect. In general, explanatory power as measured by adjusted R^2 is lower for the log percentage growth rate than for absolute growth.

C. Claim Frequency (*Claims/Payroll*)

At the marketwide level, claim frequency growth is not significantly related to rate suppression. However, this null overall effect masks a positive effect of rate suppression on claim frequency growth in the residual market and a negative but marginally significant effect in the voluntary market. These conclusions hold for both absolute growth and log percentage growth. Recall that the effect of rate suppression on reported claim frequency is theoretically ambiguous, because the positive effect on injuries, due to reduced loss control, could be offset by lower service quality including higher claim rejection rates and hence lower reported frequency. The findings here suggest that the tendency for rate suppression to increase injury rates dominates in the residual market, whereas the negative quality effect dominates in the voluntary market, as expected.

Growth in claim frequency is higher in states where the residual-market share increased over the period, although the relationship is negative but insignificant for the voluntary and residual markets separately. The positive association at the statewide level could reflect selection bias, if the proportion of premiums self-insured is positively related to growth in residual-market share and residual-

[35] Note that the larger share of small firms in the residual market could explain its higher loss growth but cannot explain the positive association between loss growth and residual-market size.

market burden; it could also reflect other omitted variables that are correlated with residual-market share growth.

The insignificant overall effect of statutory benefit growth on total losses masks different effects across submarkets. Statutory benefit growth is positively related to claim frequency growth in the residual market but negatively related to frequency growth in the voluntary market. These findings are consistent with the evidence that the cost-increasing effects of benefit growth are greater where experience rating is weak.[36]

D. Premium Growth

Premium growth per \$100 payroll is significantly positively related to rate suppression, for the total market and for the voluntary and residual submarkets separately. The coefficients for premium growth are similar for the voluntary and residual markets, whereas total loss growth is not related to rate suppression for the voluntary market. Thus the positive effect of rate suppression on voluntary-market premiums appears primarily to reflect cost shifting from the residual market, confirming that residual-market deficits are borne in part by policyholders in the voluntary market in the same state. From equation (2), for the residual market, a 10 percentage point increase in residual-market share implies an increase in absolute premium growth of .0052, about one-third as large as the sample mean premium growth of .0153. The predicted increase in log percentage premium growth for the residual market is 4.2 percentage points, compared to the mean value of 19.2. For the voluntary market, the predicted increases in absolute and log percentage premium growth from a 10 percentage point increase in residual-market share are .0044 and .039, respectively, compared to the sample means of .011 and .166. Again, these measures are clearly economically significant.

Premium growth is also significantly positively related to lagged statutory benefit growth. The coefficients for percentage growth are significantly greater than one and greater in the residual market than in the voluntary market, although coefficients for losses are less than one in all three markets. The larger coefficients in the premium equations could reflect a lag in adjusting premiums; it could also reflect other, unobservable state characteristics that are correlated with lagged benefit growth.

The statewide residual-market deficit per \$100 payroll was expected to be positively related to loss growth in the residual market, under the joint hypothesis that pooling residual-market losses undermines loss control and that this effect is greater the larger the size of the residual-market deficit. Contrary to this prediction, this variable is generally either negative or insignificant in the residual-market equations for loss growth. However, it is significantly positively related

[36] See Ruser, *supra* note 11. For claim severity, the larger and more significant effect of benefit growth in the voluntary market than in the residual market may reflect greater noise due to the small size of the residual market.

to premium growth, with a greater effect in the voluntary market than in the residual market. This is further evidence that residual-market deficits are at least partly passed through to other policyholders in the same state.

VI. Cross-Subsidy Effects of Rate Regulation

A. Theory of Political Influence

Rate regulation permits at least three types of cross subsidy: overall statewide rate inadequacy, with losses borne by insurer equity; between-class cross subsidies, which may depend on and distort the industrial mix within a state but leave unaffected the competitive positions of different firms within an industry; and within-class cross subsidies, which may affect the relative position of different firms within a class or industry. These transfers are not mutually exclusive, and the norm is probably a combination of all three. We focus here on between-class and within-class cross subsidies.

Applying the theory of political influence developed by George Stigler, Gary Becker, and Sam Peltzman, we assume that potential interest groups make rational investments in political pressure, weighing the expected benefits and costs of investing in influence.[37] Politicians and regulators are viewed as rational intermediaries, responding to influence in the form of money and votes. The political power of a group depends on the number of votes and dollar resources that it can bring to bear, which depend on the total value of the potential subsidy to the group, net of the costs of organization. Larger groups generally have higher stakes and greater voting power, other things equal. Offsetting this advantage of size is that a given subsidy rate to a large group must be financed by either a higher tax rate on other groups or a broader tax base.

For a given potential subsidy to a group, its effective power is expected to be greater the more concentrated the benefits and the smaller the number of stakeholders. This assumes that becoming informed and organizing to exert pressure entail a fixed cost per firm, such that costs of organization are larger, relative to expected benefits, the greater the number of stakeholders. A group with a larger number of stakeholders also faces greater free-rider incentives. On the other hand, it may command more votes.

The theory of political influence often does not yield sharp, testable hypotheses, and measurement of subsidies and political influence is necessarily indirect. Nevertheless, these general principles suggest several hypotheses about the characteristics of classes that are likely to be winners and losers from workers' compensation rate regulation.

[37] See Stigler, *supra* note 2; Becker, *supra* note 2; Peltzman, *supra* note 2. For application of these models to automobile insurance, see Harrington, *supra* note 7; Scott E. Harrington, Taxing Low Income Households in Pursuit of the Public Interest: The Case of Compulsory Automobile Insurance, in Insurance, Risk Management, and Public Policy: Essays in Memory of Robert I. Mehr (Sandra Gustafson & Scott E. Harrington eds. 1993). For workers' compensation, see Danzon, *supra* note 18.

B. Data and Methods

The theory of political influence suggests a general model of between-class subsidies:

$$S_{ijs}^{t} = \alpha_0 + \alpha_1 P_{ijs}^{t-1} + \alpha_2 B_s^{t-1} + \varepsilon_j + w_{ijs}^{t},$$

where S_{ijs}^{t} is a measure of subsidy to class i in industry j in state s, P_{ijs}^{t-1} is a vector of exogenous measures of political influence of the class in state s, B_s^{t-1} is a vector of state-specific controls, ε_j is a vector of fixed industry effects, and w_{ijs}^{t} is a stochastic disturbance. Our measures of subsidies and political influence and their predicted effects are as follows.

1. Measuring Subsidies

Between-Class Subsidies. We use two measures of mean class subsidy. The residual-market share of payroll (Risks) is a rough measure of percent of employees (Firms) in the class who pay subsidized rates as a result of regulation. The statewide loss ratio (Loss/Standard Premium) for the class measures the average loading charge or price of insurance for the class.[38]

Within-Class Subsidies. The hypothesis of within-class cross subsidies predicts that the level and growth of voluntary-market premiums is disproportionately related to the class-specific residual-market deficit, controlling for actuarial variables and the statewide residual-market deficit. We use the (log) voluntary premium per \$100 payroll and its growth.

2. Political Influence

Relative Class Size. Large class payroll, relative to the mean payroll per class for the state ($\text{Payroll}_{is}/\text{Payroll}_{ms}$), indicates a relatively large proportion of employees in the class and hence relatively high stakes and proportion of votes represented by the class.[39] On the other hand, a subsidy of, say, 20 percent for a class that accounts for 10 percent of state payroll requires higher taxes on other stakeholders than for a class that represents only .1 percent of state payroll. The implication is that relatively large classes may be subsidized but in any case are unlikely to be severely taxed.

High Risks. Classes with relatively high expected loss per \$100 payroll are likely to benefit from subsidies for several reasons. High-risk classes have a greater impact on an employer's overall premium cost, ceteris paribus, hence

[38] Since the standard premium for each class reflects experience rating but does not reflect dividends or schedule credits, our premium and loss ratio measures may overestimate subsidies to high risks and implicit taxes on good risks. However, insurers often deviate by a fixed percentage from manual rates, and such deviations from standard premiums are likely to be small in states with significant rate suppression.

[39] Controlling for low wages per employee (see below), large relative payroll indicates a class with a relatively large number of employees.

the potential gain from a given percentage suppression of rates is greater. Equity norms may favor cross subsidies to occupations with high loss relative to wages, which would reduce opposition to rate suppression for these classes. Subsidies to high-risk firms could yield efficiency gains by increasing the proportion of such firms that purchase coverage, thereby reducing negative externalities that may result if uninsured losses are shifted to other firms.[40] As measures of high risk, we use claim frequency (log(Claims/$100 Payroll)) and claim severity (log(Loss/Claim)). The natural log transform is used to reflect declining marginal influence of this motive for subsidy and to reduce outlier influence.

Low-Wage Workers. Actuarially fair premiums are likely to be a larger percent of payroll for low-wage occupations for several reasons. First, unskilled jobs may entail greater intrinsic injury risk, and unskilled workers may also be less skilled at avoiding accidents. Second, the WC benefit structure provides a higher wage replacement rate for low-wage workers, and the elasticity of medical loss per claim with respect to wage level is plausibly less than unity.[41] Third, since the residual-market deficit allocation is proportional to premium, the deficit tax rate per dollar of payroll and per employee is higher on classes with high premium rates per payroll dollar. This inequity of the residual-market tax system may be another reason favoring cross subsidies to low-wage classes. We do not have class-level data on wages per employee. However, loss per $100 payroll is expected to be positively correlated with low wages. Thus classes with high loss per $100 payroll are expected to be beneficiaries of rate suppression, because of both low wages and high risk. These two effects are empirically indistinguishable with our data.

Firm Size. Large firms face fewer regulatory constraints on self-insurance than small firms and have size-related advantages of risk bearing, which suggests that large firms are less likely to be taxed, other things equal. However, small firm size is likely to be correlated with in-state ownership of a high proportion of capital, which may be associated with low cost of exerting influence. In-state ownership of capital would not affect political influence if the full incidence of rate suppression—costs and benefits—is on workers, all of whom are in-state residents. But if the incidence of workers' compensation costs is partly on entrepreneurial capital or other immobile firm-specific factors, then domestic firms may be more successful at exerting political pressure than out-of-state firms. On the other hand, state-specific sunk investments are vulnerable to appropriation

[40] Coverage is optional for very small firms, many of which are high risk. Harrington, *supra* note 5, applies a similar analysis to high-risk drivers, who are major beneficiaries from rate regulation in automobile insurance.

[41] Although the income elasticity of demand for medical care is positive in broad cross sections and in time series, this does not necessarily apply to postinjury use of medical services. Moreover, because workers' compensation has very low if any co-payments and these are uniform across income levels, the insurance-induced positive correlation between income and medical care use that is observed in health insurance data (wealthier families buy more coverage and hence face lower point-of-service prices) does not apply to medical expense covered by workers' compensation.

of quasi rents, which could make in-state employer capital vulnerable to regulatory "taxes." If firms with out-of-state ownership have lower state-specific investments, they can make a credible threat to move out of state and hence are less likely to be heavily taxed.

Thus the predicted net effect of firm size on political influence is theoretically ambiguous. Small firms are expected to benefit from rate suppression under the joint hypothesis that small size is correlated with in-state ownership and in-state ownership reduces the costs of political influence by enough to outweigh the disadvantages of small size for self-insurance. As a rough measure of firm size, we use payroll per firm in the class (log(Payroll_{is}/Classcount_{is})).

Concentrated versus Diffuse Classes. The expected return to an employer from investment in influencing rates is greater for classes that account for a relatively large share of its payroll and premium. Moreover, free-rider effects are expected to reduce the political influence of classes that are dispersed over a large number of employers, compared to concentrated classes. Classes that have a relatively large number of employers that employ that class (Classcount_{is}/Classcount_{ms}) are relatively diffuse and hence are expected to be politically weak owing to free-rider effects. Classes that have a relatively large number of risks, relative to the mean number of risks per class in the state (Risk_{is}/Risks_{ms}), are the dominant class for a relatively large number of firms and hence are expected to benefit from rate suppression. However, controlling for relative payroll, a high value of relative risks could reflect predominantly small firms. In practice, high correlation between these measures makes estimation of their separate effects tentative.[42]

Class Concentration across States. Occupations that are employed in many states are likely to be relatively mobile, in which case they cannot be heavily taxed in individual states. On the other hand, classes that are concentrated in a few states may have a greater payoff and lower costs of obtaining rate suppression in those states. We use a Herfindahl index of the concentration of each class across the 27 states in our database. It is expected to be positively related to rate suppression, if the benefits of concentration outweigh the disadvantages of immobility.[43]

Industry Effects. The economic model of political influence emphasizes the importance of relative power. Thus the industry mix of beneficiary classes could differ across states, depending on industry mix. On the other hand, certain industry characteristics—for example, unionization—could enhance the political

[42] The correlation between Relative Risks and Classcount is .76; between Relative Risks and Payroll is .67; and between relative Classcount and Payroll is .96 (seven-state data).

[43] The Herfindahl index for each class is the sum of the squared shares of countrywide payroll in each state in which that class is present. It ranges from 1, for a class that is concentrated in a single state, to .034 for a highly diffuse class. The mean is .0586, indicating that most classes are relatively diffuse across the states in which the class is present. States in which a class is not present are excluded from the measure.

influence of particular classes in most states. A vector of fixed industry effects tests for such industry-specific effects.

C. Empirical Evidence of Cross Subsidies

Except where indicated, the analysis uses the same sample of 5 years for the 150 largest classes in seven states. Illinois is omitted because of missing data on Classcount. To average out some of the stochastic variation, we use 3- and 2-year mean values for periods t and $t - 1$, respectively. Lagged values are used as instruments for explanatory variables that are potentially endogenous and for political variables that are expected to affect premiums with a lag.

1. Between-Class Cross Subsidies

Residual-Market Shares. Our first test of the model of political influence uses the log odds of class-specific residual market as a measure of political influence. Table 6 reports results for both share of payroll and share of risks.[44] The first equation includes the statewide ratio of filed to approved rates, to control for the mean statewide level of rate suppression, and current and lagged statutory benefit growth, to control for possible lags in adjusting premiums to reflect benefit changes. The second equation replaces these substantive state variables with fixed state effects. The third equation adds fixed industry effects. The results are generally consistent with the theory of political influence and similar for the residual-market share of payroll and risks.

Subsidies are significantly positively related to claim frequency and claim severity, consistent with the hypothesis that classes with relatively high loss per dollar payroll, owing to either high-risk activities or low wages, have greater incentives to seek subsidies and/or face less resistance. Classes with relatively large payroll, relative to mean class payroll in a state ($\text{Payroll}_{is}/\text{Payroll}_{ms}$), have higher residual-market shares, which suggests that number of employees enhances political influence. Residual-market share is negatively related to mean firm size ($\text{Payroll}_{is}/\text{Classcount}_{is}$), which implies that small firms are more likely to benefit from rate suppression. This suggests that any advantage of large firm size, owing to lower costs of self-insurance, is on average offset by the greater political advantage of smaller firms that are presumably more frequently locally owned. The significant negative coefficient on relatively diffuse classes ($\text{Classcount}_{is}/\text{Classcount}_{ms}$) supports the hypothesis that diffuse classes are politically weak, owing to low stakes per firm and free-rider problems. The insignificant or negative coefficient on relative number of risks ($\text{Risks}_{is}/\text{Risks}_{ms}$) is inconsistent with the hypothesis that relatively dominant classes are influential. However, these conclusions are tentative because of high correlation between these variables and relative payroll size.

[44] The log odds transformation ensures that predicted values, when retransformed, are nonnegative. Results were very similar using the residual-market share as the dependent variable.

TABLE 6

Class-specific Residual Market Shares: 3-Year Means, Seven States

Variable	Share of Payroll: Equation (1)[a]	Share of Payroll: Equation (2)[a]	Share of Payroll: Equation (3)[b]	Share of Risks: Equation (1)[a]	Share of Risks: Equation (2)[a]	Share of Risks: Equation (3)[b]
Intercept	133.5519	3.8548	4.5053	89.0045	.5859	.8527
	(11.525)	(7.967)	(8.185)	(8.903)	(1.559)	(2.022)
log Claim Severity$_{is}$ $(t-1)$	.3813	.1213	.1906	.7149	.3474	.3391
	(5.604)	(1.856)	(2.428)	(11.663)	(6.581)	(5.613)
log Claim Frequency$_{is}$ $(t-1)$	.3469	.2050	.3374	.3269	.2043	.2815
	(5.381)	(3.621)	(4.805)	(6.344)	(5.298)	(5.775)
Risks$_{is}$/Risks$_{ms}$ $(t-1)$	.0101	−.0175	.0174	−.0398	−.0647	−.0449
	(.486)	(1.078)	(1.017)	(1.938)	(4.127)	(3.012)
Payroll$_{is}$/Payroll$_{ms}$ $(t-1)$	.0143	.0121	.1118	.0200	.0182	.0617
	(1.742)	(2.015)	(4.979)	(2.624)	(3.405)	(2.840)
log(Payroll$_{is}$/Classcount$_{is}$) $(t-1)$	−.4735	−.5190	−.5930	−.1371	−.1826	−.1854
	(11.748)	(13.494)	(11.473)	(4.453)	(7.114)	(5.585)
Classcount$_{is}$/Classcount$_{ms}$ $(t-1)$	. . .	. . .	−.1537	. . .	. . .	−.0714
	. . .	. . .	(5.568)	. . .	. . .	(2.662)
Statutory Benefit Growth$_s$ $(t-1)$	−8.8333	. . .	. . .	−62.1706	. . .	. . .
	(24.425)	. . .	. . .	(23.394)	. . .	. . .

Filed/Approved$_s(t-1)$	9.9565	. . .	. . .	7.8927	. . .	. . .
	(26.755)	. . .	. . .	(26.248)	. . .	. . .
Statutory Benefit Growth$_s(t)$	−6.0891	. . .	. . .	−35.6351	. . .	. . .
	(6.582)	. . .	. . .	(4.461)	. . .	. . .
Alabama	. . .	.5448	.4775	. . .	.4291	.4314
	. . .	(4.274)	(4.004)	. . .	(4.045)	(4.481)
Florida	. . .	.0433	−.1067	. . .	.5532	.4971
	. . .	(.369)	(.875)	. . .	(5.309)	(4.952)
Georgia	. . .	−.2856	−.3302	. . .	−.2111	−.2007
	. . .	(2.340)	(2.802)	. . .	(2.130)	(2.195)
Louisiana	. . .	1.5637	1.4801	. . .	1.4332	1.4202
	. . .	(11.702)	(10.942)	. . .	(13.123)	(13.589)
Maine	. . .	3.6414	3.4846	. . .	2.9476	2.9406
	. . .	(24.494)	(22.516)	. . .	(28.332)	(28.507)
Michigan	. . .	−1.2130	−1.3292	. . .	−1.4725	−1.4679
	. . .	(10.108)	(11.274)	. . .	(15.378)	(16.184)
Adjusted R^2	.6388	.7110	.7447	.6030	.7425	.7786

NOTE.—The dependent variable is $\log[R^t_{is}/(1-R^t_{is})]$. Absolute t-statistics using White standard errors are in parentheses.
[a] No industry fixed effects.
[b] Includes industry fixed effects.

TABLE 7

CLASSWIDE (Voluntary and Residual Mark) LOSS RATIO (log), 27 STATES

Variable	Equation (1): No Fixed Effects	Equation (2): Industry Fixed Effects	Equation (3): State and Industry Fixed Effects
Intercept	.1027	.1841	−1.0223
	(.859)	(1.449)	(8.805)
log Claim Severity$_{is}$ $(t-1)$	.0905	.0871	.0886
	(8.458)	(7.528)	(7.082)
log Claim Frequency$_{is}$ $(t-1)$	.1271	.1394	.1405
	(11.175)	(10.669)	(10.559)
Risks$_{is}$/Risks$_{ms}$ $(t-1)$	.2431	.0231	.0311
	(8.665)	(8.229)	(.701)
Payroll$_{is}$/Payroll$_{ms}$ $(t-1)$	−.0014	−.0017	.0904
	(1.209)	(1.469)	(.940)
log (Payroll/Risk$_{is}$) $(t-1)$	.0334	.0385	.0413
	(5.201)	(5.035)	(4.668)
Statutory Benefit Growth$_s$ $(t-1)$	.5350	.3805	. . .
	(.627)	(.443)	. . .
Herfindahl$_i$	−.2163	−.2034	−.1360
	(3.187)	(2.976)	(2.105)
(Risks$_{is}$/Risks$_{ms}$) $\times R(t-1)$	. . .	. . .	−.0084
	. . .	. . .	(.191)
(Payroll$_{is}$/Payroll$_{ms}$) $\times$ $R(t-1)$	. . .	. . .	−.0921
	. . .	. . .	(.9581)
Adjusted R^2	.0357	.0396	.1112

NOTE.—Absolute *t*-statistics using White standard errors are in parentheses.

The ratio of filed to approved rates is strongly positively related to residual-market shares, as expected.[45] Contrary to expectations, both current and lagged growth in statutory benefits are negatively related to residual-market shares. This could reflect correlation with other omitted state characteristics or reverse causality—that statutory changes to reduce benefits were enacted in states with very high residual-market shares. Of the 18 industry dummies, eight are significantly negative and the remainder are insignificant relative to agriculture (the omitted class), which suggests that agriculture is an unusually large beneficiary of rate suppression. Adding fixed industry effects increases the adjusted R^2 by only 3.3 percentage points, which suggests that industry is not an important determinant of political influence.

Classwide Loss Ratios. Table 7 reports results for the (log) classwide loss

[45] The *t*-statistics on the state variables may be upward biased owing to correlation in unobserved characteristics across classes in the same state, see Brent R. Moulton, An Illustration of a Pitfall in Estimating the Effects of Aggregate Variables on Micro Unit Data, 72 Rev. Econ. & Stat. 334 (1990). In this case, the *t*-statistics are sufficiently large that they are likely to be significant even after adjustment for potential bias.

ratio (voluntary plus residual market).[46] The sample in these equations includes all classes in the 27 states for which 5 years of NCCI data are available. The first equation includes only class-specific variables; the second equation adds fixed industry effects; the third equation adds state fixed effects. To test whether the measures of relative influence have greater impact in heavily regulated states, we include two interactions with an indicator *R*, defined as one for states in the top third of the distribution of residual-market shares and zero otherwise. The low adjusted R^2 reflects the extreme variation in these class-specific loss ratios. Nevertheless, some variables are significant and consistent with predictions.

Loss ratios are significantly positively related to lagged claim frequency and claim severity, which implies that classes with relatively high loss/payroll or relatively low wages per employee are systematically subsidized, consistent with the evidence from residual-market analysis in Table 6. Loss ratios are positively related to firm size ($\log(\text{Payroll}_{is}/\text{Risks}_{is})$), which implies that large firms pay lower prices. This plausibly reflects scale economies in expense loadings for large firms[47] and hence is not inconsistent with the evidence in Table 6 that residual-market shares are inversely related to firm size. Loss ratios are significantly higher for classes with a relatively large number of risks, consistent with the prediction that dominant classes benefit from regulation. Effects are not significantly different in heavily regulated states.

The coefficient of the Herfindahl index is negative. The implication that more diffuse classes have higher loss ratios suggests that mobility across states protects a class against excessive rates. Adding fixed industry effects has a negligible effect on the adjusted R^2, which confirms that industry effects are not strongly consistent across states. By contrast, state fixed effects explain an additional 7 percentage points.

2. Within-Class Cross Subsidies

In Table 8, both voluntary-market premium levels and premium growth are significantly positively related to the class-specific residual-market deficit, in addition to the statewide deficit. This is consistent with the hypothesis of within-class as well as between-class cross subsidies. Voluntary-market premiums are significantly lower for classes with relatively large payroll and significantly higher for diffuse classes, after controlling for actuarial factors. This is consistent with earlier results that classes with relatively large stakes or voting power and relatively concentrated classes are more likely to benefit from cross subsidies. Premium levels are also lower for classes that are frequently dominant classes for employers. The actuarial variables are as expected: premiums and premium growth are positively related to expected losses (lagged $\text{Loss}_{is}/\text{Payroll}_{is}$) and

[46] The log transform reduces the influence of extreme values. Observations at more than 4 standard deviations from the mean were eliminated.

[47] Chelius & Kavanaugh, *supra* note 11.

TABLE 8

CLASS-SPECIFIC VOLUNTARY-MARKET PREMIUM LEVELS AND GROWTH, SEVEN STATES

	log(Premium/Payroll)[a]		GROWTH IN PREMIUM/PAYROLL[a,b]	
VARIABLE	Equation (1)	Equation (2)	Equation (1)	Equation (2)
Intercept	6.6868	9.2859	.1435	.1785
	(2.249)	(3.291)	(1.318)	(1.694)
log ($Loss_{is}$/$Payroll_{is}$) ($t-1$)	.4951	.4374	.0077	.0066
	(13.819)	(12.111)	(7.412)	(5.819)
log ($Payroll_{is}$/$Classcount_{is}$) ($t-1$)	−.1139	−.1154	−.0017	−.0006
	(6.318)	(6.434)	(2.925)	(.852)
$Risks_{is}$/$Risks_{ms}$ ($t-1$)	−.1168	−.0924	−.0003	−.000295
	(7.679)	(7.619)	(.786)	(1.059)
$Classcount_{is}$/$Classcount_{ms}$ ($t-1$)	.0668	.0336	−.0010	−.0009
	(3.313)	(2.222)	(2.491)	(2.361)
$Payroll_{is}$/$Payroll_{ms}$ ($t-1$)	−.0593	−.0281	.0012	.0009
	(3.535)	(2.294)	(3.209)	(2.792)
Statewide $Deficit_s$ ($t-1$)	69.5119	113.9467	7.6868	8.2567
	(1.495)	(2.647)	(5.063)	(5.487)
Class $Deficit_{is}$ ($t-1$)	17.6029	20.8575	.8472	1.372273
	(2.545)	(2.165)	(2.395)	(5.326)
Statutory Benefit $Growth_s$ ($t-1$)	−.9041	−1.2214	.0661	.0565
	(1.790)	(2.453)	(3.911)	(3.599)
Statutory Benefit $Growth_s$ (t)	−5.5287	−7.8164	−.1526	−.1943
	(2.178)	(3.140)	(1.752)	(2.214)
Adjusted R^2	.6914	.7255	.2749	.2844

NOTE.—Absolute t-statistics using White standard errors are in parentheses.
[a] Industry fixed effects in second equation only.
[b] $(\text{Premium/Payroll})^t - (\text{Premium/Payroll})^{t-1}$.

negatively related to firm size ($Payroll_{is}$/$Classcount_{is}$), which confirms that large firms pay lower expense loadings. Industry fixed effects again add little additional explanatory power.

VII. CONCLUSIONS

These class data for workers' compensation insurance confirm the basic hypothesis that rate suppression exacerbates loss growth and hence ultimately leads to higher premiums, contrary to its initial intent. These effects are economically as well as statistically significant. The loss-increasing effects are greater in the residual market than in the voluntary market, consistent with greater employer moral hazard in the residual market, due to greater suppression of overall rate levels, attenuation of experience rating, and possibly weaker insurer incentives for loss control in the residual market. The finding of no significant relation between loss growth and the size of the statewide residual-market deficit is inconsistent with the joint hypothesis that pooling reduces insurer incentives for loss control and that the effect becomes greater as the total size of the residual market and deficit increase.

There is strong evidence that rate suppression in workers' compensation results in persistent cross subsidies at three levels: between classes within a state, between low- and high-risk employers within classes in a state, and to policyholders statewide, presumably from insurer equity. The analysis here has focused on between- and within-class cross subsidies. These increase as statewide rate suppression increases (see the Appendix). The patterns of cross subsidies are consistent with predictions from a simple model of political influence. High risks and possibly low-wage workers are generally beneficiaries, owing to higher stakes and possibly less opposition. Relatively large classes and relatively concentrated classes generally benefit from rate suppression. Small firms are beneficiaries, presumably because of advantages in local influence or because small employers have greater personal stakes in influencing outcomes owing to immobile capital. Industry effects are generally weak, but state-specific effects are strong. Understanding the forces that lead to severe statewide rate suppression in some states remains an important topic for future research.

Purely mechanical regulatory lag may contribute to the delay in adjusting premiums in the face of exogenous loss shocks. However, a simple regulatory lag model cannot readily explain how rate inadequacy can continue for more than 5 years, spanning multiple rating periods and despite clear evidence of the inadequacy. The simple regulatory lag model also cannot explain the evidence on redistribution, with some of the cost of rate inadequacy shifted to voluntary-market policyholders. Moreover, the pattern of redistribution between classes is consistent with predictions of a simple model of political influence. Since rate suppression is counterproductive as well as obvious, its persistence seems implausible in the absence of political pressure from beneficiaries.

The evidence here suggests that rate regulation permits short-term gains by some groups, notably those with high risks, at the expense of those with lower risks and insurer equity. But in the long run it is a negative-sum game because rate suppression exacerbates loss growth. There may be a social concern to subsidize high-risk firms and activities, although no such objective is explicit in the regulatory language. This analysis suggests that the costs of such subsidies should be weighed against any perceived benefits. Rate regulation that encourages risky behavior and hence increases the growth of workers' compensation losses is a high-cost mechanism for achieving these goals.

APPENDIX

Further Evidence of Cross Subsidies

As a test of the hypothesis that at least some of the increase in statewide mean rate suppression is borne by other policyholders within the state, we examined the relationship between mean loss ratio and dispersion of loss ratios across classes within a state. If an increase in statewide rate suppression entails increased cross subsidies between classes within the state, then states with relatively high mean loss ratios should have greater dispersion in loss ratios between classes, reflecting the use of excessive premiums (relative to losses) for taxed classes to subsidize inadequate premiums for other classes. If a higher mean loss ratio for a state overall is not associated with greater between-class dispersion in loss ratios, then the increase in subsidies to some classes must be borne primarily by insurer equity or by policyholders in other lines.

We identified states in the top and bottom quartiles of the distribution of state mean loss ratios in the earliest 2 policy years (period B) and in the most recent 3 policy years (period A). For each of these states, we estimate the distribution of loss ratios across classes. The evidence confirms that states with a higher mean statewide loss ratio have a higher interquartile range. For states in quartile 4 in period A, the interquartile range across classes ranges from 2.4 to 5.7, with a median of .90. For states in quartile 1 in period A, the range is from 1.7 to 4.0, with a median of .60. A similar pattern appears in period B. This evidence supports the hypothesis that at least some of the deficit increase associated with increased statewide mean rate suppression is borne by taxed classes in the same state.

To test the hypothesis that statewide rate suppression effects cross subsidies between industries within a state, we performed a similar analysis at the level of the industry. If firms in the same industry have common class structures and lower costs of coordinating lobbying efforts, for example, through use of industrywide union and employer organizations, then cross subsidies should follow industry lines. We find little evidence that cross subsidies follow industry lines, at least using the NCCI definition of industries, which suggests that industrywide lobbying is undermined either by differences in the class mix across firms in an industry or by other obstacles to cooperation.

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